Downward Nominal Wage Rigidity in Italy: Evidence and Consequences

by
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Abstract:

This paper uses administrative longitudinal micro-data from the Italian Social Security Institute (INPS) to estimate the extent of downward nominal wage rigidity. The determinants of wage changes are explicitly modelled, as is the measurement error deriving from the fact that earnings inclusive of benefits, not hourly wages, are available in the data. Estimates show that the degree of downward nominal wage rigidity is medium/high – between 51% and 68% of all notional wage cuts being prevented by the existence of proportional rigidity. The implications of the estimated nominal wage rigidity for the real side of the economy are also explored.

Keywords: Nominal wage rigidity, measurement error, proportional and threshold rigidity models, natural unemployment rate.

JEL classification: E24, E31, J31

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1. Introduction

The existence of wage rigidities has traditionally occupied, and even more often preoccupied, entire generations of economists, of any school and inspiration. Wages that do not duly respond to the changing economic conditions of the markets are held to be rigid, or more optimistically, sticky. Market adjustment processes - unable to rely on the re-equilibrating forces of prices - are then bound to unburden themselves on something else, quantities above all.

As macroeconomics textbook predicate, nominal wage rigidity is one of the possible reasons for the existence of an upward sloping aggregate supply, and the resulting short-run trade-off between unemployment and inflation – the known Phillips curve – at the root of the Keynesian stabilization policies. Other economists point instead to real wage rigidities – and the institutional aspects (unions but also firms' strategic behaviour in the presence of market imperfections) responsible for wage levels incompatible with the labour market equilibrium – as the main cause of high and persistent levels of unemployment rates.

In Italy the vast territorial differences have sometimes lead to believe that an increased flexibility in the wage bargaining mechanisms may contribute to reinstate the convergence processes between the rich and developed north and the poorer and backward south of the country.

The inflation targets that monetary authorities should pursue have also been suggested to depend on the extent of wage rigidity. For, if on the face of downwardly rigid nominal wages it may be desirable to accept a positive rate of inflation (which would cool down real wage growth and "grease" the wheels of the economy), it would instead be optimal to aim at zero inflation if wages are perfectly flexible. This way, in fact, the efficiency costs spread in the economic system by wrong price signals (which throw "sand" on the economy wheels) may be avoided (Akerlof et al., 1996).

But if wage rigidity occupies such a central place in economic theory and policy, a question which logically should precede, and therefore inform, any relating debate should be: Are wages really rigid? And how much? Since it is difficult to reconcile the hypothesis of wage rigidity with the postulate of rational behaviour (e.g., Romer, 1996 – put in references)[1], it is particularly important to know whether the validity of the hypothesis can be confirmed empirically.

Surprising as it might seem, finding a satisfactory answer to such a legitimate question is not easy enterprise, for Italy as well as for many other industrialised countries. Empirical research has of course not failed to quantify the relationship at a country level between wage growth (or price inflation) and unemployment rates, estimating Phillips curves and wage adjustment processes over time. However, these estimates have mainly been carried out at an aggregate level (time series), distant by its own nature from the optimising behavioural processes of the single economic agents (firms and employees)) and unable to fully recognise the heterogeneity of such agents. Because the presence of inertia in wage adjustments ultimately derives from the optimal choices of rational agents (under various forms of information, institutional and technological constraints), a microeconomic approach appear to be at least as desirable as the traditional macro perspective employed in early research on the topic. The recent acquisition of longitudinal micro-datasets – particularly absent in Italy - has contributed to extending the scope for micro-oriented research on wage rigidity.

[1] Some model builders, for instance, reject DNWR on the grounds that it implies money illusion of the economic agents.

A burgeoning literature has already provided evidence of the existence of some downward nominal wage rigidity (henceforth, DNWR) in a number of countries (notably the US, UK, Germany, Canada, Switzerland), but no such empirical investigation has been undertaken in Italy as yet.[2] The aim of this paper is to contribute to this literature by studying the Italian case, drawing on recently available administrative data from the Italian Institute for Social Security (INPS). First, the overall degree of nominal wage rigidity will be estimates using various econometric techniques. Then, the estimates will be used to derive the real implications of the observed dynamics of wage changes, in terms of their costs for the long-run unemployment rate and output.

Assessing the extent and the effects of nominal rigidities is all the more important in a country which has been standing out for its relatively high and persistent unemployment rate throughout the eighties and the nineties – at least when compared to other EU countries. The presence of fairly strict employment legislation protection in the Italian labour market provides additional reasons for investigating the extent of nominal wage rigidity. For, as Bewley (2000) suggests, it may be possible (and indeed this is what he finds for the US) that layoffs are a tool preferred by managers to nominal wage cuts when in need to restore their firm's competitiveness, as these cuts would have an adverse effect on employees' morale. While the available data are all but ideal to

[2] A related literature on the estimation of the so-called wage curve has empirically investigated the dynamics of *real* wages around their equilibrium level and the real wage rigidities at the root of the equilibrium ("natural" or long-run) unemployment rate (e.g., Layard *et al.*, 1991). The dynamics and rigidity of real wages in Italy have been explored by Lucifora and Orrigo (1999). As the consequences of nominal wage (or price) rigidity are often strengthen by the presence of real rigidities, the present paper (and related literature) constitute an essential element for a complete assessment of the macroeconomics policy trade-offs. Note that real rigidities (e.g., due to unions or efficiency wage considerations) – while potentially implying the existence of a positive natural rate of unemployment – need not entail a short run trade-off between inflation and unemployment (e.g., Romer, 1996, ch. 6).

confront Bewely's prediction to Italian practice, the modelling strategy can at least provide some preliminary evidence on the matter.

The paper is organised as follows. Section 2 briefly describes wage-setting practises in Italy, which have important implications for both our modelling strategy and results' interpretation. Section 3 provides details about the data used and the sample selection, while section 4 describes the empirical evidence on the distribution of wage changes. In section 5 the modelling approach is explained, while results are illustrated in section 6. Section 7 then explores the real consequences of the estimated degree of nominal wage rigidity. Some final considerations are gathered in section 7.

2. Wage Setting Practices in Italy

Italy's wage setting process has been and still is dominated by industry wide national unions' wage contracts. These are formally binding only for the firms belonging to the employers' associations who have signed them. However, both the courts (in case of worker-firm disputes) and the law (which subordinates some firms' benefits to the use of those contracts and computes employers' social security contributions as a % to be applied to the maximum between actual earnings and those fixed by national contracts) somehow extend their actual coverage. Therefore the non-covered sector may be identified with the "hidden" economy, estimated to be around the 15% of total employment in 1998 by ISTAT (the national statistical office).[3]

Firms' level bargaining is quite widespread in at least the largest firms, which are however a minority in Italy's landscape. The contracts there signed top up national wages and, particularly in the periods of stronger unions' power (since the mid '60s to

[3] To a large extent, and particularly in the less developed and high unemployment *Mezzogiorno*, the hidden economy features pertain not only to the avoidance of the tax and social security contributions burden but also to the use of sub-standard wages and working relations rules.

the beginning of the ‘80s), “anticipated” the issues subsequently generalised through industry wide contracts. Wage rises dictated at the firm level through unions’ bargaining are quite egalitarian (within the firm itself).

An opposite nature have the autonomous firms’ wage policies. Both these and individual worker-firm bargaining had been quite compressed during the period of stronger unions’ power. However, since the mid ‘80s these components have acquired some role. Presently, these components represent between 5 and 10% of average earnings, another 10% being dictated by firms’ level contracts.

The Italian pyramidal wage bargaining (whereby wage contracts can be bargained at the national, industry and firm level, with agreements struck at a higher level immediately becoming lower bounds for bargaining at lower levels) limits the scope for downward wage flexibility at the firm level. This has to be borne in mind when, say, comparing the Italian experience with that of the US, where instead bargaining is mainly conducted at the firm level. Employment protection is stronger too, particularly for large firms, which suggests that firms in Italy will have to find alternative ways to circumvent the labour market inflexibility (both in terms of wages and quantities), for example through cautious intertemporal smoothing of wage and employment changes, or through the strategic use of more flexible components of wages such as benefits and overtime.

As far as nominal adjustment to price inflation is concerned, important changes have taken place during the ‘90s. Up to 1992 nominal wages were indexed to prices through an automatic mechanism known as *scala mobile* whose rules were uniform across sectors. Actually, given the relevance of such an automatic mechanism, the above mentioned predominance of industry wide bargaining may be questioned. From the end of WWII to 1975 price inflation triggered each quarter wage rises differentiated across sectors and job categories[4]. The adjustment was asymmetric, as it operated only in case of price rises, and was meant to safeguard a given minimum threshold real wage (differentiated across sectors and job categories)[5]. Over time, as real wages were rising the actual safeguard granted by the mechanism declined and upward adjustments of the threshold wage levels were agreed upon several times at national level. In 1975 one of these periodical re-adjustments took place. On top of incorporating the large real wage rises contracted for during the previous years, the agreement profoundly changed the general rules: the quarterly wage rises were made uniform across sectors and job categories lifting up the fixed amount wage rises granted to the lowest paid sectors and job categories. This increased the average degree of safeguard against price inflation (which peaked at around 100% in 1978-1979) and imparted a strong egalitarian bias to the overall mechanism (low paid workers were automatically gaining ground in case of price inflation).

The scala mobile quite soon came under attack for its inflationary bias. The high degree of safeguard provided for was a source of real wage resistance against terms of trade shocks (particularly the oil prices' hikes experienced in 1974 and 1979) and indirect taxes rises. The quarterly timing speeded up the inflationary spiral. The egalitarian bias was affecting wage differentials and restricting the role for autonomous firms' and unions' decisions as the scala mobile automatisms were responsible for most of the wage dynamics.

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[4] Up to the end of the '60s the wage rises were also differentiated across regions as the national contracts themselves, while agreed upon at the national level, provided for wage categories differentiated across regions (the so called gabbie salariali).

[5] On average the actual safeguard against price rises (defined as the % increase in nominal wages triggered by a .01% price rise) was around 50% in mid '70s.

In 1986 – the first year available in the sample used below for the empirical analysis is 1985 – the mechanism was partially reformed. Its timing became half-yearly. Both the average safeguard granted for and the egalitarian bias were reduced: a 100% safeguard was granted to a minimum uniform wage threshold, with a 25% safeguard granted to the difference between the nationally contracted for wage (differentiated across industries and broad job categories) and that common minimum threshold.[6] On average the safeguard against price rises declined to around 60%.

In 1992 the scala mobile was finally dismantled. The formal agreements signed up in July 1992 and July 1993 depicted a new bargaining system, in which national contracts, to be agreed upon every two years (against the 3 years of the previous set up[7]), are supposed to be guided by the price inflation expected (and targeted by the Government) for the future, while firms' level bargaining is supposed to be geared by profit sharing considerations. Past inflation triggers automatic wage rises only in case no agreement is reached, and the safeguard granted is at most 50%. The difference between actual and targeted inflation is not automatically recovered: it has to be taken into account when a new bargaining deal starts. In such a case, however, unions and employers' associations have also to consider the reasons for such a discrepancy, in particular taking into account whether it was caused by terms of trade shocks against which employees are not anymore automatically insured.

3. Data, Definitions and Sample Selection

For the analysis of nominal wage rigidity in Italy I use administrative data from the Italian Institute for Social Security (INPS), containing information on a sample of

[6] No automatic safeguard was granted for the topping up of wages bargained for at the firm or at the individual level.
[7] In the current set up national contracts deal with more regulatory aspects every 4 years.

employees over a period of twelve years – from 1985 up to 1996. For each calendar year 1985/1996, the Social Security forms of employees born on 10 March, June, September and December of any year were selected. In this way a sequence of random samples of the population of employees is formed (sampling ratio 1:91). Each yearly sample includes approximately 100,000 employees of Italian private firms, with the exclusion of workers in agriculture and the central state administration. Using available identifiers (fiscal and social security codes), individual longitudinal data can be generated for each worker. The firm's longitudinal records are then accessed for each worker in the sample and the employer's attributes are linked to the employee, obtaining a matched employer-employee database. The data therefore include not only individuals' wage and career histories but also a certain number of characteristics of each worker and of the firm where s/he currently works and has held previous jobs. Among personal characteristics, we have information about the employee's gender, age, geographical region where s/he was employed, along with his/her job qualification. Information about the firm where the job is held includes the firm's sector of activity, dates of opening and closure and its occupational trend (number of manual and non-manual workers, along with total remuneration to the two occupational groups).

As we do not observe the actual number of hours worked by an employee, we cannot compute hourly wages. We do however know the number of "paid days"[8] of each employee in each job spells and his/her total remuneration (inclusive of overtime and fringe benefits), which allows us to compute the employee's daily wage. It is the year-to-year changes in this nominal wage rate that lie at the heart of my examination of downward rigidity.

[8] Some of these paid days may not be worked days, since, included under this heading are periods of paid time during which no work is done, e.g., maternity leave, sick leave, holiday.

Since the domain of the data source is limited to employment in the private sector, another limitation of the data is that periods between employment spells cannot be characterized in a precise way. In other words, it is not possible to distinguish between individuals' movements to unemployment, self-employment, employment in the public sector and exits from the labour force. Moreover, the data do not identify the reasons for a job separation, thus preventing us form distinguishing between employees' layoffs and quits.

The sample of workers whose wage changes are put under observation has been restricted so as to reflect a high labour market attachment. In particular, the paper focuses on full-time year-round workers, aged between 15 and 64 and who are job stayers (keep a job in the same firm in both years in which the wage is compared). Descriptive statistics are collected in Appendix A, Table 2.

I trim the sample by excluding 1% of wage change observations in each tail. The effect of the trim is to exclude wage changes that involve cuts of greater than 23% and increases of greater than 38%, which might be thought of as resulting from mis-reporting of either total remuneration or worked days or both. To further minimize the measurement error deriving from variability in the amount of work, I have also performed all computations after restricting the sample to all employees working exactly 312 days in the year (equivalent to 52 weeks a year). These are workers in regular and stable jobs and it might be held that estimated wage dynamics and nominal rigidities is higher than in the overall labour market, where more precarious and flexible jobs also exist. The results obtained with this restricted sample (which I will refer to as "stable" workers) are then compared to those obtained over the sample of employees reporting *any* number of worked and paid days in the year (what I will refer to as "all workers”). Finally, estimates are also presented for the subsample of men only, which allows to investigate if the dynamics of nominal wages is different across gender. A quick inspection to part A (all workers), B (“stable” workers) and C (“stable” men) of Table 2 reveals that the three samples are very different in terms of characteristics. About 55% of all workers report exactly 312 paid days per year and, compared to all workers, tend to have slightly higher mean wage growth, to be employed in larger and expanding firms, to be older, more concentrated in the north of the country, more likely to be white collars and male, and in older firms. When further restricting to men employees only, these differences – which are not unexpected - further enlarge, but remain overall fairly small, and unlikely to produce major differences in the estimates of nominal wage rigidity.

4. The Distribution Of Wage Changes: Descriptive Evidence

The distribution of nominal wage changes in Italy does not show up the typical regularities that have been found for Britain and the US. For example, while big spikes at zero wage growth are often reported – and been referred to as a clear hint of downward nominal wage rigidity – no such thing is perceptible in our data. For each year in the sample, Figure 1 and Table 1 illustrate the distributions of one-year wage changes, measured in terms of log wage differences. A vertical line at zero wage growth has been drawn in Figure 1, as well as a line indicating the value of inflation for the year considered.

Besides the absence of any sizeable mass of concentration at zero growth, three important features are noticeable in Figure 1. First, the distributions appear to be centred on a positive value of wage growth, which is generally equal or above the inflation rate registered for the year. The last two columns of Table 1 show that inflation was for example about 5.8% in 1985, while the median wage change was around 6.7% (and the median wage growth was 7.6%).[9] As a result, more than half of the employees in 1985 saw their real wages increase, perhaps reflecting a period of productivity growth. Inflation in subsequent years was not very different, reaching its highest value in 1989 at about 6.5% and its minimum at the end of the period, in 1996, when prices grew on average by 4%. The central location of the distributions depicted in Figure 1 in general follows the behaviour of inflation over time (see also Figure 3B), though the two do not always move in the same direction in the short run. In particular a compression of real wages can be detected during the years 1991-1994 when the median wage change is pushed below the level of inflation by the economic slowdown that characterized those years (see Table 1). In fact, while before 1991 median wage growth always exceed inflation by at least 1 percentage point, the two become almost indistinguishable thereafter.

A second distinctive feature of the wage change distribution is that numerous negative changes co-exist with a majority of wage increases, resulting in distributions that – though fairly skewed to the right in some years – look generally and surprisingly more symmetric than one would expect. As reported in Table 1, in 1985 about 8% of wage changes were wage cuts, with a median drop of almost 4%. In contrast, 89% of employees experiencing changes in their annual pay saw their nominal wage increase. The proportion of wage cuts reaches its highest value in 1992, at the heart of the period of economic slowdown. Correspondingly, the proportion of wage increases fall (at 82%) and the median raise dropped at only 5.7%. Figure 3A suggests that the proportions of

[9] Mean wage growth is always higher than the median, and in 1985 was at 8.7%.

wage cuts and freezes are inversely related to inflation, while figure 3B visualizes the positive correlation existing between inflation and mean wage changes.

The third characteristic of the distributions of nominal wage changes is the low percentage of employees whose wage was unaltered over two consecutive years (wage freezes). In fact, on average about 3% of the sample experienced zero wage growth. This result contrasts to what has been reported for the US and UK, where spikes at zero are clearly visible in each year. For instance, Smith (2000), reports that "…on average 9.0% of non-job changers experienced exactly zero annual growth in their usual gross weekly pay during 1992-6". Even higher were the corresponding figures found for the US (see, Card and Hyslop, 1996).

Taken at their face value, the results discussed in this section would suggest that the structure of nominal wage changes is compatible with a labour market characterised by a fairly high degree of flexibility, as it would appear that there are no big constraints for firms to reduce the wages they pay if they so need do. However, this description may be a surprising one for Italy, traditionally included in the list of countries with rigid labour market institutions. The possibility that our simple descriptive evidence is unable to reveal the "true" – and much higher – degree of downward nominal wage rigidity is investigated in the following section. An econometric approach is employed, aiming at estimating the extent of "frictions" in the labour markets, controlling for (1) various observable determinants of the underlying rigidity-free wage changes, and (2) the likely presence of measurement errors in the available data. The key question I attempt to address with the econometric model is whether the relatively high proportion of wage cuts revealed by the preliminary data inspection is real or reflects instead errors in measured wages. In particular, in my sample wage changes may result from variation in the amount of work supplied (overtime, hours worked), strategic mis-reporting of remuneration or days worked by firms[10], fridge benefits as well as traditional measurement error. For example, daily wages are computed by dividing total remuneration by the number of days each employee is paid for. But paid days may not coincide with worked days, as the former include period of pay time during which no work is done (maternity leave, sick leave, holiday). The econometric methodology applied below aims at circumventing our inability to observe these wage components separately in the data. The same methodology has been followed for studying nominal wage rigidity in Germany (Knoppik and Beissenger, 2001), the US (Altonji and Devereux, 1999), Switzerland (Fehr and Goette, 2000), Canada (Fares and Seamus, 2000) who have all shown the importance of controlling for measurement error in estimating the true degree of nominal wage rigidity.

5. Econometric methodology

Following the notation of Knoppik and Beissinger (2001), I call *notional wage change* that change in the nominal wage that would have occurred in the absence of nominal downward wage rigidity. For individual *i* at time *t* this is denoted by Δw_{it}^{*} and is assumed to be explained by a vector X_{it} of characteristics as follows:

$$\text{(1)} \qquad \Delta w_{it}^{*} = X_{it}\beta + \varepsilon_{it}$$

where β is a vector of parameters to be estimated and $\varepsilon_{it} \sim N(0,\sigma_{\varepsilon}^{2})$ represents the effect of i.i.d normally distributed error terms.

The *actual wage change* Δw_{it}^{a} is equal to the notional one, whenever the latter is positive. If the notional wage change is negative, it is assumed that the actual change

[10] That Italian firms – particularly in the south of the country – might find it convenient to underreport the actual number of days worked by the employee has been suggested by Contini et al. (2000).

will be negative too when the individual is not affected by DNWR; if instead the individual is affected by DNWR, then the observed wage change will be zero. Whether the individual's notional wage is constrained by DNWR is unobservable, and is described by a random variable D_{it} - taking value 1 if individual *i* at time *t* experiences DNWR, and 0 otherwise, with probabilities:

(2) $\quad$ Prob(D_{it} =1) = ρ $\quad$ and $\quad$ Prob(D_{it} =0) = 1- ρ.

The behaviour of nominal wage changes is then described by the following model:

$$(3) \qquad \Delta w_{it}^{a} = \begin{cases} X_{it}\beta + \varepsilon_{it} & \text{if } \ 0 \le X_{it}\beta + \varepsilon_{it} \\ 0 & \text{if } \ X_{it}\beta + \varepsilon_{it} < 0 \ \text{ and } \ D_{it} = 1 \\ X_{it}\beta + \varepsilon_{it} & \text{if } \ X_{it}\beta + \varepsilon_{it} < 0 \ \text{ and } \ D_{it} = 0 \end{cases}$$

This is a model with proportional downward wage rigidity, since a proportion ρ of notional wage cuts will not occur due to the presence of rigidity. The degree of rigidity is captured by the parameters ρ, which can be estimated by he data.

When estimating (3) one complication arises from our inability to observe hourly wages, as our data contain only information on changes in daily wages (earnings), which are denoted by Δy_{it}. Observable changes in earnings can be interpreted as being the sum of changes in actual wage rates and a random variable, μ_{it}, which can capture (unobservable) variation in working hours, fridge benefits, bonuses and more conventional measurement errors:

$$(4) \qquad \Delta y_{it} = \Delta w_{it}^{a} + \mu_{it}$$

Together with equation (3), equation (4) implies that:

$$(5) \qquad \Delta y_{it} = \begin{cases} X_{it}\beta + \varepsilon_{it} + \mu_{it} & \text{if } \ 0 \le X_{it}\beta + \varepsilon_{it} \\ \mu_{it} & \text{if } \ X_{it}\beta + \varepsilon_{it} < 0 \ \text{ and } \ D_{it} = 1 \\ X_{it}\beta + \varepsilon_{it} + \mu_{it} & \text{if } \ X_{it}\beta + \varepsilon_{it} < 0 \ \text{ and } \ D_{it} = 0 \end{cases}$$

Various distributional assumptions can be adopted for the measurement error μ_{it}, each implying a different degree of correctness in the measurement of the dependent variable. According to (6), all observations in the sample are measured with errors, and it is assumed that these are distributed normally. I will refer to the resulting model as a normal measurement error model (NME).

(6) $$\mu_{it} \sim N(0, \sigma_{\mu}^{2})$$

The working of the models in (3) and (5)-(6) is intuitively illustrated in picture 2. Consider first model (3) where measurement error is assumed away. In this case look at the bold lines of the figure only and neglect the sparse points in the graph. Any negative notional change results in a corresponding actual wage change after it is being allocated to either of the two branches in the left part of figure 1. Given the absence of measurement error, observed wage freezes automatically imply that the notional wage cut has been forced to lie on the horizontal branch in the left part of the picture. Given the determinants of the notional wage change, the percentage of wage freezes in the data is therefore the crucial information to estimate ρ. Now consider the model in (5)-(6). In this case, the presence of measurement error results in the sparse points depicted in the graphs: a notional wage cut which is not implemented due to the DNWR constraint need not lye on the horizontal axis anymore. In this case, even with no observed wage freezes in the sample, the model will allocate some of the observed wage cuts to the horizontal branch and some to the 45-degree line brunch, the best fitting resulting in the estimate of ρ.

In the literature of nominal rigidity more general versions of the models specified above have also been estimated, notably by assuming that only a fraction of the observations be affected by measurement errors. For example, one can suppose that that only a proportion 1-p of the observations are measured with (normally distributed) errors, while for the remaining observations μ_{it} is equal to zero, obtaining a mixed measurement error model (see Altonji and Devereux, 1999; Beissenger and Knoppik, 2001). These formulations may be attractive as they do not suffer from one potentially important shortcoming of the models with normal measurement error (and p=0), namely that they are inconsistent with a spike in wage change at 0. However, the benefit of using these more general models may be less striking in our data as such spike is not observed. Estimates obtained with these more general formulations are nonetheless presented in what follows, with all result tables collected in Appendix A.

Appendix B derives the likelihood functions implied by each of the models described in this paper, which are maximised to obtain the estimates of the parameters of interest.

6. Estimation Results

The determinants of wage changes

The set of explanatory variables entering the notional wage equation (1) includes:

(i) age and age squared (scaled variables, /100)

(ii) occupation dummies: white collar, manager, apprentice (reference category: blue collar)

(iii) industry dummies, denoted sector1-7 (reference category: manufacturing, food/“alimentary”, textile)

(iv) dummies for firms that between year *t* and *t*+1 have increased the number of employees (“growing”), or reduced (“shrinking”).

(v) the (log of) firm size and its square.

(vi) inflation and lagged inflation

(vii) national unemployment and its lagged values

(viii) a dummy for female, and its interaction with age

(ix) regional dummies for North West, North East, Centre and South

(x) dummies for the age of the firm

(xi) a time trend

Similar variables have also been used by the existing papers on downward nominal wage rigidity, and differences mainly arises form the data availability (e.g., most papers enter education as an explanatory variable, but the INPS data do not contain it). In all the models that I have estimated, the impact of observed heterogeneity on wage changes (the X variables described above) is generally highly significant and of the expected sign. As we are not particularly interested in the effects of these variables – the focus being on estimating the amount of rigidity existing in the system – I will briefly comment them with reference to an initial OLS equation describing the determinants of wage changes, which does not control for the effect of measurement error nor does it estimate the amount of frictions in the system (see Table 3). All models have been estimated for (i) the sample of all full-time workers with at least one worked day per each month in the year ("all workers" sample), (ii) the sample of workers for whom exactly 312 paid dais have been reported ("stable workers" sample), and (iii) the sample of men workers with exactly 312 days ("stable men" sample). The results for the three sub-samples did never differ sensibly, so I will not discuss them in detail. One interesting implication is that – while female wage dynamics is a bit different than male's, as shown below – this has not strong implications for the estimate of nominal wage rigidity, particularly so once the sample of employees in stable jobs has been selected. This is not surprising, as the literature has long suggested that the main gender differences are to be found in the participation decision processes, while conditioning on employment – and even more so on stable employment – remaining differences are less dramatic.

Estimation also took account of the fact that the sample data consisted of observations on individuals from the same firm at each *t*, and repeated observations on the same individuals across successive pairs of periods, because I pooled wage changes from my panel. These repeated observations mean that the i.i.d. assumption is violated. To account for this, I use a pseudo maximum likelihood estimator, as follows. The complex survey statistics literature has developed methods for adjusting the estimates of the parameter of the covariance matrix to account for sample clustering, using formulae that allow for arbitrary correlation between observation within the same cluster. See *inter alia* Huber (1967), Binder (1983) and White (1982). I defined each cluster to consist of all the employees who were members of the same firm at year *t* (which very often means that the employee is in the same firm in successive pairs of years too). The sample log-likelihood is a 'pseudo-likelihood' in this case (Gourieroux and Monfort, 1996), from which can be derived a 'robust' variance estimator of the parameter estimates using Taylor series linearization. None of the paper on nominal wage flexibility that I know of produce robust standard error, though their data are likely to present similar clustering problems.

In the OLS regressions of Table 3, individual's age is found to have a negative impact on wage changes, consistently with the "classic" profile of wage levels with respect to age, which is found to be increasing at decreasing rates. Age squared is here intended to capture the presence of higher order polynomials in age. White collars and managers tend to have higher wage growth than blue collars. Though a bit more surprisingly, the same seems to be true for apprentices.

Female employees have a lower wage growth than male's, but tend to reduce the distance with their male counterpart as they grow older and acquire labour market experience, as shown by the interaction between the dummy for female and age.

Expanding firms often grant wage increases, as reflected by a positive coefficient of the dummy "growing"; the opposite occurs for "shrinking" firms. Those working in large firms too seem to obtain bigger wage increases than in smaller firms. There is also some evidence than employees in firms aged less than 5 years, or more than 10, manage to get larger wage raises than those working in the reference category (firm aged 5-10 years).

As expected, nominal wages are highly responsive to inflation: when this grows, wages grow too and workers get protected by various institutional arrangements – such as the *scala mobile* in operation until 1992 – as well as by re-negotiations aimed at maintaining unaltered the real purchasing power of their wages. Strangely enough, though, past inflation enters the notional wage change equation with a negative sign.

Consistently with a typical Phillips-curve argument, the rate of unemployment has a negative coefficient, indicating that when there are many unemployed wages tend to decline. The difference between current and lagged unemployment (a proxy for the deviation of current unemployment from its equilibrium level) has a negative coefficient, as it is generally found in the literature.

A negative time trend is supported by the data, capturing - in a crude way as it is - the effect of various institutional changes occurred in the labour market over the time period considered, notably the slow but constant phasing out of the *scala mobile*, as described in section 2.

Estimated downward nominal rigidity

As about the estimated degree of rigidity, Table 4 displays the results obtained with the model (3), which does not control for measurement error. For the three sub-sample considered, the estimated degree of rigidity is fairly low, with ρ equal to about 21%, implying that less than one out of four wage cuts are actually prevented by the existence of any rigidity in the system. This should not come as a surprise. Given the low percentage of wage changes that are exactly zero, and the relatively high share of wage cuts (see Table 1) that, more importantly still, are interpreted by model (3) as "true" changes, rather than the result of measurement error, it might have been expected that the structure in (3) can fit the data only by making the estimated ρ fairly small.

Once again, one could then be tempted to draw the conclusion that the system at study hardly displays any rigidity at all. However, when starting worrying about the impact of measurement errors things soon change. For example, when we recode the observations with very low wage changes (between -.5% and +.5%) as exact zeros, and re-estimate the same model, then get a higher and higher value of ρ. Instead of arbitrarily deciding which small wage cuts are to be regarded as exact zero wage changes, the models I estimate below explicitly introduce measurement errors, thereby allowing for the possibility that some of the underlying zero changes may actually display themselves as non-zero changes.

The estimates of the model with proportional rigidity described in (5) when measurement error is assumed to be normal as in (6) are reported the results in Table 4. The effect of observed characteristics is not very different from what we have seen in the previous model, and will not be further discussed. The standard deviation of the error term entering the notional wage change equation (1) is at about 0.07. The standard deviation of the measurement error term μ is estimated at about 0.03 and is significant, confirming the importance of allowing for an imperfectly observed dependent variable in the model. In Table 4 it is reported a sizeable increase in the estimated value of ρ, which is now at about 61-63%, depending on the sub-sample considered.

Many of the observed wage cuts are treated by model (5) as arising from measurement errors, rather than being "true" adjustments. On the other hand, some of these observed cuts are now identified as would-be wage freezes which, due to measurement error, do not manifest themselves in the form of exact zero wage growth. As a result, the estimated ρ is now much higher and implies that six in ten notional wage cuts are not implemented due to the operation of downward nominal wage rigidity.

Table 5 displays the estimates obtained when model (5) is generalised to include mixed measurement errors, whereby a proportion p of the observations is assumed to be exactly measured, μ=0, while the remaining 1- p are measured imperfectly and the error term is assumed to be normally distributed. In this case, p is estimated to be between 59 and 65% (depending on the sub-sample considered), while the estimate of the degree of rigidity is only slightly different than before: ρ now lies in the range 51% and 68%.[11] What is interesting to note in this model formulation, is that the reduction of the proportion of exactly measured observations (that in the models in Table 4 was by definition equal to zero) is accompanied by an increase in the variance of the measurement error variable μ_{it}. This makes sense and implies that the model is trading-off the number of observation subject to measurement error with the size of the error variability: fewer observations report measurement error in model (5) than in model (3), but with a higher variance.

Before moving to briefly analyse a few alternative models (next section) and the implication of the estimated rigidity for the real side of the economy (section 7), it is worth spending a few words on a potential trade-off between downward wage rigidity and employment flexibility. In fact, one might pose that where employment protection is strict, firm's may attempt to re-gain some degree flexibility - in the face of shocks in the demand for their output – by making wage cuts more likely and loosening the constraint of DNWR. This would also be in line with Bewely's predictions that managers prefer workers' dismissal – when this is an open option - to the adverse effect of wage cuts.

In this respect, note that the sample of "all workers" and that of "stable workers" differ in the important dimension that the second selects all workers in regular and guaranteed jobs, while included in the first are also those employees in more precarious

[11] Values of ln$L$ are not comparable between the models with normal measurement error and those with mixed measurement error because of the mass point in the latter.

and flexible jobs. Firms have therefore less room for employment dismissal in the second sample. Accordingly, one might expect that this disadvantage is traded-off by firms in terms of higher wage flexibility represented by a lower estimated ρ. This is what is indeed found in Table 4, as ρ is equal to 0.63 in the sample of "stable workers" and 0.61 for "all workers". However, this ranking is completely reversed by the estimates of the model with mixed measurement error in Table 5, preventing us from any definitive conclusions about the existence of the trade-off.

Some Alternative Models

One way in which the models presented in the previous section can be extended is to allow the rigidity parameter ρ to depend on observed characteristics. None of the existing papers on nominal rigidity have implemented such extensions. One reason might be that, though ease to implement, the resulting models can be difficult to estimate (the number of parameters increases and the data might be not rich enough to identify them). Likelihood convergence problems indicated that this is indeed the case in my data. Here I will only display the results obtained when ρ in the model with normal measurement error was allowed to depend on firm's size and year dummies (other variables were considered as well, but they were either not significant or made the model difficult to converge). Two interesting results emerged. First, ρ tends to be larger for smaller firms. This might perhaps be surprising, but one can propose a sensible explanation. Large firms face a stricter employment protection legislation than small firms do.[12] Accordingly, the negative relation of firm's size with the extent of

[12] For instance, the debated article 18 of the Workers' Statute establish that employees can only be dismissed with evidence of "fair cause" in firms with at least 15 employyes, while the same clause does not applies for firms with less than 15 employees.

downward nominal rigidity may suggest that large firms may recover some degree of general flexibility on their ability to cut nominal wages. On the contrary, small firms – freer to fire workers – are required to offer more implicit guarantee of wage protection. When a dummy for firms with more than 15 employees was entered in the equation for ρ, its coefficient was estimated to be negative and significant. The estimates in Table 8 display a different specification of the relationship between the extent of nominal rigidity and firm's size, by using a quadratic in (log) firm's size. In this case, the result is less neat: up to firm size of about 65 employees, the relationship is negative, but then turns positive for firms of medium to high employment.

As for the dependence of ρ and year dummies, the results in Table 8 would suggest that over time the degree of rigidity has increased.[13] However, this result was not robust. Indeed, separate regression by years always failed to converge. Presumably, the reason has to do with the very small numbers of zero or negative wage changes that can be used in yearly regression to estimate ρ (number which is instead sufficiently higher when pooling the sample). A look at figure 1 reveals that later years have a fatter tails on the left of zero wage growth, i.e. there more observations to estimate ρ in the most recent years in the sample. This might explain the sign of the year dummies in Table 8, and cast doubts on the evidence that rigidity has increased over time.

In the literature on wage flexibility models like (5), with proportional downward wage rigidity, are not the only types that have been estimated. Certain studies (Altonji and Devereux, 1999; Fehr and Goette, 2000) have indeed preferred models with *threshold* rigidity, where two additional parameters are introduced and estimated.

[13] The same result was obtained if including a linear time trend in the equation for ρ.

According to this approach, small cuts in the notional wage (up to a cut equal to $\alpha>0$) do not generate changes in observed wages, save for those attributable to measurement errors. If however the cut in the notional wage is big enough (larger than α), then observed wages will drop as well but by λ% ($\lambda>0$) less than what suggested by the behaviour of the notional wage change (see figure 3).

$$(7) \qquad \Delta y_{it} = \begin{cases} X_{it}\beta + \varepsilon_{it} + \mu_{it} & \text{if} \quad 0 \le X_{it}\beta + \varepsilon_{it} \\ \mu_{it} & \text{if} \quad -\alpha \le X_{it}\beta + \varepsilon_{it} < 0 \\ X_{it}\beta + \varepsilon_{it} + \mu_{it} + \lambda & \text{if} \quad X_{it}\beta + \varepsilon_{it} < -\alpha \end{cases}$$

Note that this wage model contains as special cases both a model of perfect wage flexibility and a model of perfect downward nominal wage rigidity. In the case of perfect flexibility, both α and λ are zero. For perfect downward nominal wage rigidity, λ is arbitrary and α approaches ∞. Because these models are nested in the general model (7), one can test whether the restrictions implied by either perfect rigidity or perfect flexibility are consistent with the data (see Altonji and Devereux, 1999).

Table 6 shows the estimates of the threshold rigidity model described by (7). Once again the impact of observable characteristics is of the same magnitude and sign as in the models presented earlier. The variances of the two error terms too do not display any sizeable change with respect to the model with proportional rigidity.

As for the rigidity parameters, it is found that all intended wage cuts that do not overtaken a threshold α equal to 7% are not implemented as a result of the constraints erected by this form of downward wage rigidity (e.g., menu costs), and any small wage cut which is not exactly zero is then attributed to measurement error. If, instead, intended wage cuts are larger than α, then observed wage changes are negative too, and of the amount indicated by the notional wage, as λ was estimated to be virtually zero (and constrained to be so in the models displayed in Table 7). Once again, this evidence points to a labour market where complete upward flexibility has to be set against the presence of various institutional constraints that limit the ability of firms to reduce nominal wages by the amount they deem appropriate given the economic environment.[14] This approach to the estimation of nominal wage rigidity has however not be further pursued in the paper for the reason explained in the next session.

7 Implications of the estimated nominal rigidity

A number of measures have been used by the literature to provide a hint of the real implication of the estimated downward nominal rigidity. One advantage of the proportional rigidity model over the threshold rigidity model is their ability to provide a direct estimate of the proportion of employees affected by DNWR, which can then be used to provide implications for the costs of these rigidity. Besides, there are also contingent reasons for preferring the former approach: the estimates provided are fairly robust to various specification and sample selections (as shown in section 6), while convergence problems for the threshold model indicated that this approach is a less appropriate description of the nominal wage dynamics and rigidities for Italy during the sample period. The estimated consequences of nominal rigidity that I discuss below therefore use the estimates obtained with the proportional rigidity model in (5). The alternative measurement assumptions for the measurement error (normal or mixed) made little difference in the estimated implications of nominal rigidity.

The first measure that I consider is the share of the observations in the sample which can be expected to actually face the constraint represented by DNWR. Though the

[14] Estimates obtained by Altonji and Devereux (1999) using model (7) on their PSID sample are as follows: α=.32, λ=.17, σ^2_μ=.0014 and σ^2_ε=.017, suggesting that their data show more downward rigidity and less measurement error problems (see their Table A2).

Italian labour market seems to show a high degree of resistance to implementing nominal wage cuts when these cuts are required, one in fact can ask if the constraints are actually binding for any large number of individuals. For it is not sufficient to know that, in principle, drops in notional wage changes may be prevented by the operation of nominal rigidity; one has also to assess how often the actual individual, firm and aggregate conditions suggest that a wage reduction (a drop in the notional wage) is required. If notional wage cuts are very unusual – given the prevailing conditions – then it may not matter a lot that these cuts would be prevented by a fairly high degree of downward nominal rigidity. In fact, one can write the probability, r_{it}, of a wage freeze for individual *i* at time *t* as:

$$(8) \qquad r_{it} = \hat{\rho} \cdot \Pr\left(\Delta w_{it}^{*} < 0\right) = \hat{\rho} \cdot \Phi\left(-x_{it}\hat{\beta} / \hat{\sigma}_{\varepsilon}\right).$$

The share of observations *R* affected by DNWR is then estimated by:

$$(9) \quad R = \frac{1}{N}\sum_{t}\sum_{i} r_{it}$$

where *N* is the number of observations in the sample.

When the estimates of Table 4 are used, I compute that $\hat{R}$=11% only. This confirms the idea that, though nominal wages can be considered as fairly rigid downwardly, the conditions of Italian labour market during the eighties and the nineties have generally required wage increases, particularly important to keep up with inflation and productivity growth.

The measure *R* can also be computed by postulating various scenarios of steady-state level of inflations. When the economy is in steady state there are no expectations of changes in inflation and unemployment. A simple way of examining how binding it is the constraint of DNWR for various levels of steady-state levels of inflation, I have recomputed r_{it} in (8) after:

(i) setting current and lagged changes in inflation to zero,

(ii) replacing observed inflation with various assumptions about the level of steady-state inflation, and

(iii) force the coefficient of inflation to be equal 1 (as postulated by a standard steady-state formulation of the Phillips curve). After making these adjustments, the resulting share of observations affected by DNWR in the sample – as a function of a postulated level of inflation – is computed as in (9) and is denoted $R(\pi)$.

Figure 5 depicts the values of $R(\pi)$ for different values of π. It shows that the impact of DNWR is relatively high for very low rate of steady-state inflations, with about 18% of the observations in the sample receiving wage freezes if inflation is equal to zero. As inflation grows, R gets smaller; at an inflation rate of, say, 10% less than 3% of the observations in the sample would be affected by DNWR.

Another measure typically computed by the literature to quantify the real relevance of the estimated DNWR, is the extent by which the expected observed wage change is higher than the expected notional change. In fact, when DNWR is in place, a certain number of notional wage cuts are not implemented, which implies that the expected notional wage growth is less than the expected observed wage growth. The difference between these two expected changes is called the sweep-up, which is computed at an individual level as:

$$\text{(10)} \qquad \text{sweep}_{it} = E\Delta w^{a}_{it} - E\Delta w^{*}_{it}$$

$$= \hat{\rho} \cdot \left[\hat{\sigma}_{\varepsilon}\phi\left(x_{it}\hat{\beta}/\hat{\sigma}_{\varepsilon}\right) + x_{it}\hat{\beta}\Phi\left(x_{it}\hat{\beta}/\hat{\sigma}_{\varepsilon}\right) - x_{it}\hat{\beta}\right]$$

where $E\Delta w^{*}_{it}$ is – from (1) - simply $x_{it}\hat{\beta}$, while $E\Delta w^{a}_{it}$ can be computed from (5). Aggregate sweep-up is obtained by averaging over the sample's observations:

$$(11) \qquad AS = \frac{1}{N}\sum_{t}\sum_{i}\text{sweep}_{it}$$

Using the estimates of Table (5), the aggregate sweep up in my sample was about 0.005. This means that DNWR has the effect of increasing the expected observed wage growth over the notional by about 0.5 percentage points. One can also compute the aggregate sweep-up for various scenarios of steady-state inflations, as done with respect to $R(\pi)$. The results indicate that at zero inflation, the sweep up is about 1% point, and gets smaller as inflation raises, up to becoming negligible for inflation rates of say 10%.

Finally, one can also compute measures that assess the effect of DNWR on the long-run unemployment rate. According to a standard accelerationist Phillips curve, inflation will accelerate or decelerate depending on whether unemployment is below or above the *natural rate*, while any existing rate of inflation will continue if unemployment is at the natural rate. The natural rate is thus the minimum, and only, sustainable rate of unemployment, but the inflation rate is left as a choice variable for policymakers. Since complete price stability has attractive features, many commentators who accept the natural rate hypothesis believe the central bank should target zero inflation. Akerlof *et al.* (1996) question the standard version of the natural rate model and each of these implications. They investigate the consequences of accounting for DNWR in a model that otherwise resembles a standard natural rate model and show that there is no natural unemployment rate. Rather, the rate of unemployment that is consistent with steady inflation itself depend on the inflation rate. In the long run, a moderate steady rate of inflation permits maximum employment and output in the simulated version of their model. Maintenance of zero inflation, instead, measurably increases the sustainable unemployment rate and correspondingly reduces the level of output.

Central to their argument, is a modified version of the Phillips curve, which they write as follows:

(13) $\pi_t = \pi_t^e + a(u^{LS} - u_t) + s_t$

where π_t^e denotes the expected rate of inflation, u_t is the rate of unemployment in t, u^{LS} is the lowest sustainable rate of unemployment and s_t is a term reflecting the effect of DNWR on the standard accelerationist Phillips curve. In particular, s_t is interpreted as a shift in expected unit labour costs arising from DNWR and enters linearly in (13) the same as a shift in labour costs arising from any other reasons different than DNWR. In the definition of Akerlof et al., s_t is the *real wage wedge* relative to the level of the real wage (RWW), which measures the wedge (and therefore the cost) introduced by DNWR between the expected aggregate actual and notional real wage levels. It can be easily shown that, in turn, the RWW is equal to the aggregate sweep up, AS.[15] In the long run, when $\pi_t = \pi_t^e$, equation (13) implicates that the unemployment rate with non-accelerating inflation, the so-called NAIRU, is written as:

(14) $u^{NAIRU} = u^{LS} + \frac{1}{a} s_t$

The NAIRU can be larger than the lowest sustainable rate of unemployment if the relative real wage wedge s is positive. DNWR – by creating a wedge between the actual and notional real wage – can indeed create an excess long-run unemployment (u^{NAIRU-}

[15] In fact, RWW= $E\left[\frac{1}{N}\sum_t\sum_i \mathrm{w}_{it}^{a} - \frac{1}{N}\sum_t\sum_i \mathrm{w}_{it}^{*}\right] = \frac{1}{N}\left[\sum_t\sum_i \mathrm{Ew}_{it}^{a} - \sum_t\sum_i \mathrm{Ew}_{it}^{*}\right] =$
$\frac{1}{N}\left[\sum_t\sum_i \mathrm{E}(\mathrm{w}_{it}^{a} - \mathrm{w}_{it-1}^{a}) - \mathrm{E}(\mathrm{w}_{it}^{*} - \mathrm{w}_{it-1}^{*})\right] = SU$

u^{LS}) given by (1/*a*) s_t, where s_t is of the same size as in (11) and can again be computed for different level of steady-state inflation. The only missing information to compute the excess long-run unemployment in (14) is an estimate of *a*. The values reported for Germany (Knoppik and Beissenger, 2001) and the US (Stiglitz, 1999) range from about a=0.1 to about a=0.5. Rather than using any of the available estimates for Italy (e.g. Golinelli, 1998), I have here preferred to report the excess long-run unemployment rate obtained for various values of *a* and for alternative scenario for the rate of inflation, including values (see Table 9). At zero inflation, a value of alpha equal to 0.4 (as the one used by Knoppik and Beissenger for Germany) implies that DNWR produces additional long-run unemployment of about 2 percentage points. This shows the danger of aiming at a very low inflation rate in a country whose labour market institutions and practise still entail a relatively high set of constraints to firm's desired wage changes. As estimates for Germany, a country that shares many of the Italian labour market rigidities, are in the range of those reported here, EU central bankers have additional reasons to think twice before setting extremely low inflation targets or either should strongly encourage the so-often invoked EU labour markets reforms.

7. Conclusions

Using administrative longitudinal micro-data from the Social Security Institute (INPS), this paper has estimated the extent of downward nominal wage rigidity in Italy. The descriptive analysis of the year-to-year wage change distribution reveals little degree of downward nominal wage rigidity, with many of the annual distributions revealing that no important spikes at zero wage growth are observed, while wage cuts seem to occur even in expansionary periods. The tentative conclusion that Italy's wage structure is flexible is however at variance with the conventional description of the country as with fairly rigid labour market institutions.

The empirical strategy of the paper has then been that of using an econometric model apt to circumvent the limits of the data. The determinants of wage changes have been explicitly modelled, as well as the measurement error deriving from the fact that earnings, not hourly wages, are observed in the data, and earnings are gross of such items as benefits and overtime. Proportional and threshold rigidity models have been estimated, both pointing to the existence of significant impediments to the firm ability to implement optimal wage reductions. According to the first approach, around 50%-64% of all notional wage cuts are prevented by the existence of proportional downward nominal wage rigidity. With the second approach, a threshold of about 7% has to be overtaken before any desired wage cut can be implemented. On the contrary, models of rigidity that did not account for measurement error delivered much lower estimates of the degree of rigidity, and distorted the real picture of the wage dynamics almost as much as the descriptive evidence of the wage change distributions.

The estimated models were then used to compute various implications of the extent of downward nominal wage rigidity, particularly in terms of its costs for the long-run unemployment rate. Though moderate, these costs are not negligible and might suggest that, given the existing degree of downward nominal wage rigidity, a zero inflation policy might be a costly option, while a low but non-zero target might deliver a lower long-run unemployment rate and higher output.

A note of caution should however accompany the conclusions of the paper. It has been argued that a labour market with downwardly rigid nominal wages can bring about important macroeconomic costs in the face of a low inflation rate, as real wages would be made sticky by both nominal wage rigidities and slow moving prices. In this respect, the case of Italy falls somewhat in between, as a high degree of nominal rigidities is accompanied – since the mid eighties through the mid nineties – by a rate of inflation that is better described as moderate, rather than low. It is unclear, though, how far results – and in particular their real consequences - obtained with the levels of inflations of the eighties and nineties can be carried over to scenarios where prices grows by, say, only 2%.

Finally, future research may add further important insight on a number of important issues that the present paper has been unable to investigate, such as the assessment of how far firms supplement wages with bonuses or use overtime as a means of circumventing nominal wage rigidity. The availability of richer dataset is crucial in this respect.

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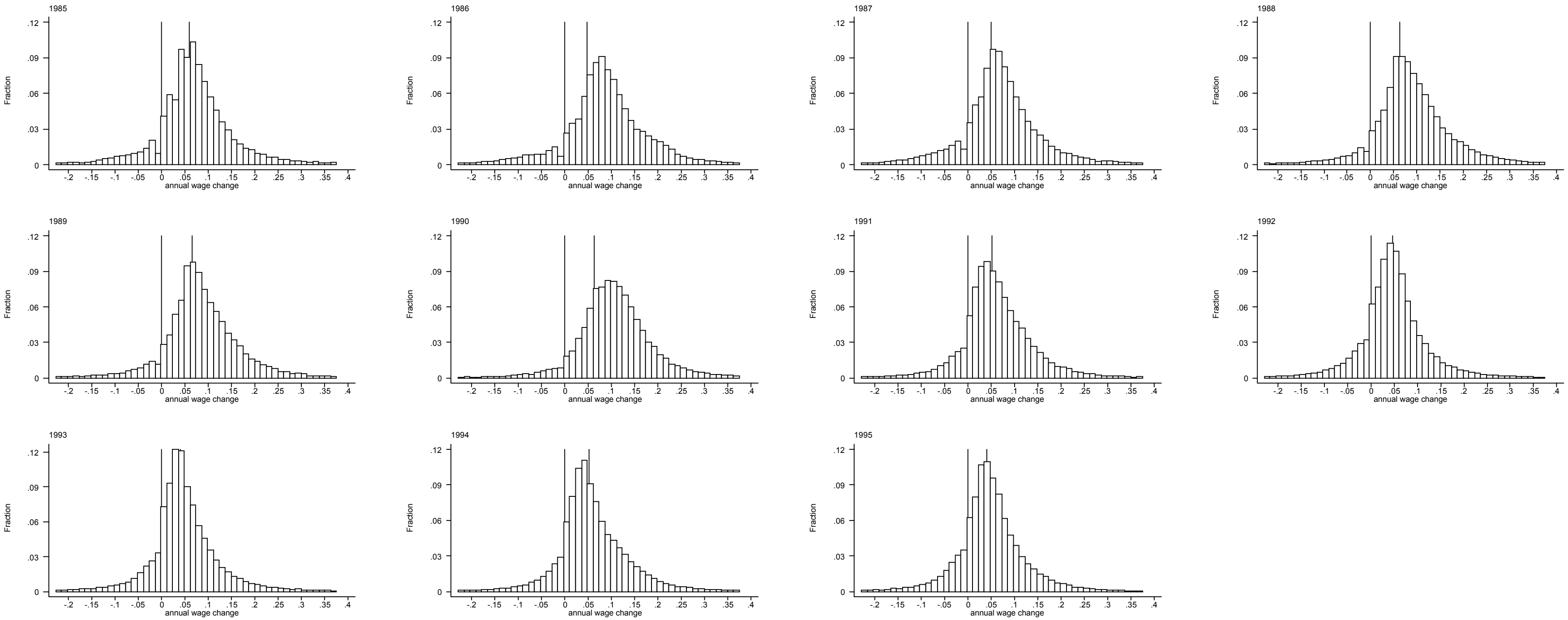
Appendix A: Tables and Figures

Table 1: Wage changes and inflation

year	Wage freezes	wage cuts	wage increase	Number of Observations	median wage cut	median wage increase	Median Wage change	inflation rate
1985	0.04	0.08	0.89	16742	-0.038	0.076	0.067	0.058
1986	0.02	0.05	0.94	22077	-0.043	0.097	0.093	0.047
1987	0.02	0.08	0.90	23669	-0.037	0.078	0.071	0.051
1988	0.02	0.06	0.92	24652	-0.037	0.090	0.084	0.063
1989	0.02	0.06	0.92	20814	-0.038	0.087	0.081	0.065
1990	0.01	0.04	0.95	20579	-0.041	0.112	0.107	0.063
1991	0.03	0.11	0.86	24595	-0.034	0.069	0.058	0.053
1992	0.03	0.15	0.82	23810	-0.034	0.057	0.048	0.046
1993	0.04	0.14	0.82	24992	-0.035	0.051	0.041	0.041
1994	0.03	0.11	0.86	26103	-0.033	0.061	0.052	0.052
1995	0.03	0.15	0.81	26552	-0.033	0.056	0.044	0.040
total	0.03	0.09	0.88	254585	-0.037	0.076	0.066	0.053

Notes: Wages changes exceeding the outside the 1st-99th percentile range have been trimmed. Then the sample has been restricted to employees working 312 days a year.

Figure 1: Wage change distributions, 1985-1996. All full-time year-round workers



Note: All workers sample. Two vertical bars are drawn for each distribution: one at zero wage growth, the other at the inflation rate for the year.

Table 2: Descriptive statistics

A) All workers

variable	Obs	Mean	Std. Dev .	Min	Max
ln(wage change)	462307	0.069	0.079	-0.228	0.375
ln(firm size)	461644	5.00	2.917	0	11.65
ln(firm size) squared	461644	33.55	35.10	0	135.68
worker's age	462307	0.037	0.010	0.015	0.063
age squared	462307	1.507	0.777	0.225	3.97
growing firm	462307	0.415	0.493	0	1
shrinking firm	462307	0.405	0.491	0	1
sector1	461861	0.036	0.187	0	1
sector2	461861	0.095	0.294	0	1
sector3	461861	0.258	0.437	0	1
sector5	461861	0.165	0.371	0	1
sector6	461861	0.074	0.262	0	1
sector7	461861	0.114	0.318	0	1
north East	462162	0.234	0.424	0	1
centrer	462162	0.187	0.389	0	1
south	462162	0.163	0.369	0	1
female	462301	0.275	0.446	0	1
female*age	462301	0.009	0.016	0	0.063
manager	462065	0.009	0.096	0	1
apprentice	462065	0.013	0.112	0	1
white-collar	462065	0.387	0.487	0	1
white collar * age	462065	0.014	0.019	0	0.063
inflation	462307	0.052	0.008	0.039	0.064
lagged inflation	462307	0.056	0.0124	0.040	0.091
unemployment rate					
level	462307	1.020	0.098	0.86	1.16
first difference	462307	0.280	0.693	-1.1	1.3
lagged first differ.	462307	0.318	0.710	-1.1	1.3
firm's age					
< 1 year	461857	0.033	0.177	0	1
1-5 years	461857	0.116	0.320	0	1
10-20 years	461857	0.387	0.487	0	1
> 20 years	461857	0.322	0.467	0	1
trend	462307	6.150	3.158	1	11

Legend.
sector1: "energia, gas, acqua"; sector2: "estratt., manif. trasform. min."; sector3: "manif. trasformaz. metal."; sector4 (omitted category): "manif. aliment., tess."; sector5 "commercio, ingr., dett."; sector6 "trasporti e comunic."; sector7: "credito., assicuraz., servizi impr."

B) "Stable" workers (reporting 312 paid days a year)

variable	Obs	Mean	Std. Dev .	Min	Max
ln(wage change)	254585	0.073	0.073	-0.228	0.375
ln(firm size)	254189	5.173	2.934	0	11.65
ln(firm size) squared	254189	35.37	35.98	0	135.68
worker's age	254585	0.038	0.010	0.015	0.063
age squared	254585	1.544	0.769	0.225	3.969
growing firm	254585	0.448	0.497	0	1
shrinking firm	254585	0.380	0.485	0	1
sector1	254290	0.052	0.223	0	1
sector2	254290	0.092	0.289	0	1
sector3	254290	0.219	0.414	0	1
sector5	254290	0.198	0.398	0	1
sector6	254290	0.076	0.266	0	1
sector7	254290	0.159	0.366	0	1
north East	254497	0.250	0.433	0	1
centrer	254497	0.178	0.382	0	1
south	254497	0.132	0.339	0	1
female	254582	0.265	0.441	0	1
female*age	254582	0.009	0.016	0	0.063
manager	254370	0.014	0.119	0	1
apprentice	254370	0.005	0.070	0	1
white-collar	254370	0.535	0.498	0	1
white collar * age	254370	0.020	0.020	0	0.063
inflation	254585	0.052	0.008	0.039	0.065
lagged inflation	254585	0.056	0.012	0.040	0.092
Unemployment rate					
Level	254585	1.022	0.099	0.86	1.16
first difference	254585	0.270	0.684	-1.1	1.3
lagged first differ.	254585	0.315	0.702	-1.1	1.3
firm's age					
< 1 year	254289	0.023	0.165	0	1
1-5 years	254289	0.109	0.312	0	1
10-20 years	254289	0.380	0.485	0	1
> 20 years	254289	0.350	0.477	0	1
Trend	254585	6.279	3.140	1	11

C) "stable" men (with 312 paid days per year)

Variable	Obs	Mean	Std. Dev .	Min	Max
ln(wage change)	187081	0.074	0.075	-0.228	0.375
ln(firm size)	186813	5.45	2.957	0	11.65
ln(firm size) squared	186813	38.48	37.47	0	135.68
worker's age	187081	0.039	0.010	0.015	0.063
age squared	187081	1.631	0.776	0.225	3.969
growing firm	187081	0.453	0.498	0	1
shrinking firm	187081	0.395	0.489	0	1
sector1	186876	0.064	0.244	0	1
sector2	186876	0.101	0.302	0	1
sector3	186876	0.241	0.428	0	1
sector5	186876	0.174	0.379	0	1
sector6	186876	0.093	0.289	0	1
sector7	186876	0.142	0.349	0	1
north East	187016	0.247	0.431	0	1
centrer	187016	0.181	0.385	0	1
south	187016	0.150	0.357	0	1
female	187081	0	0	0	0
female*age	187081	0	0	0	0
manager	186949	0.0182	0.134	0	1
apprentice	186949	0.004	0.063	0	1
white-collar	186949	0.470	0.499	0	1
white collar * age	186949	0.019	0.020	0	0.063
inflation	187081	0.052	0.008	0.039	0.065
lagged inflation	187081	0.056	0.012	0.040	0.092
unemployment rate					
level	187081	1.021	0.099	0.86	1.16
first difference	187081	0.267	0.684	-1.1	1.3
lagged first differ.	187081	0.313	0.702	-1.1	1.3
firm's age					
< 1 year	186875	0.026	0.158	0	1
1-5 years	186875	0.099	0.293	0	1
10-20 years	186875	0.380	0.485	0	1
> 20 years	186875	0.371	0.483	0	1
trend	187081	6.236	3.134	1	11

Table 3. Wage change regression. OLS estimates.

Rigidity type	None (simple OLS)		None (simple OLS)		None (simple OLS)	
Measurement error	No		No		No	
Sample	all workers		stable workers		all stabel men	
dip. Var. : ln(wage change)	Coeff.	robust s.e.	Coeff.	robust s.e.	Coeff.	robust s.e.
ln(firm size)	0.002	0.000	0.000	0.000	0.001	0.000
ln(firm size) squared	0.000	0.000	0.000	0.000	0.000	0.000
worker's age	-1.347	0.089	-1.460	0.116	-1.466	0.133
age squared	0.013	0.001	0.013	0.001	0.013	0.002
growing firm	0.005	0.001	0.006	0.001	0.006	0.001
shrinking firm	-0.004	0.001	-0.003	0.001	-0.003	0.001
sector1	0.016	0.002	0.012	0.001	0.011	0.001
sector2	0.005	0.001	0.005	0.001	0.004	0.001
sector3	0.003	0.000	0.003	0.001	0.003	0.001
sector5	0.004	0.000	0.000	0.000	-0.001	0.001
sector6	0.007	0.001	0.002	0.002	0.001	0.002
sector7	0.008	0.001	0.005	0.001	0.003	0.001
north East	0.002	0.000	0.000	0.000	0.000	0.000
centrer	-0.002	0.000	-0.001	0.001	-0.002	0.001
south	-0.004	0.001	-0.004	0.001	-0.005	0.001
female	-0.016	0.001	-0.019	0.001		
female*age	0.265	0.026	0.338	0.032		
manager	0.025	0.001	0.024	0.001	0.024	0.001
apprentice	0.076	0.001	0.079	0.002	0.071	0.003
white-collar	0.031	0.001	0.026	0.001	0.026	0.001
white collar * age	-0.463	0.028	-0.396	0.029	-0.404	0.035
inflation	0.730	0.080	0.519	0.089	0.564	0.106
lagged inflation	-0.089	0.065	-0.172	0.042	-0.162	0.049
unemployment rate						
level	-0.037	0.011	-0.042	0.016	-0.037	0.020
first difference	-0.011	0.000	-0.013	0.001	-0.013	0.001
lagged first differ.	0.003	0.001	0.004	0.002	0.005	0.002
firm's age						
< 1 year	0.003	0.001	0.004	0.001	0.004	0.001
1-5 years	0.001	0.001	0.001	0.001	0.001	0.001
10-20 years	-0.001	0.000	0.000	0.001	0.001	0.001
> 20 years	-0.001	0.001	-0.001	0.001	-0.001	0.001
trend	-0.002	0.000	-0.003	0.000	-0.003	0.000
constant	0.108	0.014	0.148	0.019	0.141	0.024
sigma	0.076		0.070		0.072	
Number of observations	461247		253882		186615	
Number of clusters	64327		41663		29455	
Log-Likelihood						
R-squared	0.074		0.090		0.087	

Table 4. ML estimates of wage changes and downward wage rigidity.

Rigidity type	proportional		proportional		proportional	
Measurement error	No		No		No	
Sample	all workers		stable workers		stable men	
dip. Var. : ln(wage change)	Coeff.	robust s.e.	Coeff.	robust s.e.	Coeff.	robust s.e.
ln(firm size)	0.002	0.000	0.000	0.000	0.001	0.000
ln(firm size) squared	0.000	0.000	0.000	0.000	0.000	0.000
worker's age	-1.327	0.091	-1.451	0.119	-1.462	0.136
age squared	0.013	0.001	0.013	0.001	0.013	0.002
growing firm	0.005	0.001	0.006	0.001	0.006	0.001
shrinking firm	-0.004	0.001	-0.003	0.001	-0.003	0.001
sector1	0.016	0.002	0.012	0.001	0.012	0.001
sector2	0.005	0.001	0.005	0.001	0.004	0.001
sector3	0.003	0.000	0.004	0.001	0.003	0.001
sector5	0.004	0.000	0.000	0.000	0.000	0.001
sector6	0.007	0.001	0.001	0.002	0.001	0.002
sector7	0.008	0.001	0.004	0.001	0.003	0.001
north East	0.002	0.000	0.000	0.000	0.000	0.000
centrer	-0.002	0.000	-0.001	0.001	-0.002	0.001
south	-0.004	0.001	-0.005	0.001	-0.006	0.001
female	-0.016	0.001	-0.019	0.001		
female*age	0.259	0.026	0.335	0.033		
manager	0.025	0.001	0.025	0.001	0.024	0.001
apprentice	0.077	0.001	0.080	0.003	0.071	0.003
white-collar	0.031	0.001	0.027	0.001	0.027	0.002
white collar * age	-0.462	0.029	-0.393	0.029	-0.400	0.036
inflation	0.741	0.081	0.527	0.091	0.573	0.107
lagged inflation	-0.096	0.067	-0.184	0.044	-0.172	0.051
unemployment rate						
level	-0.038	0.011	-0.043	0.017	-0.038	0.020
first difference	-0.011	0.000	-0.013	0.001	-0.014	0.001
lagged first differ.	0.003	0.001	0.004	0.002	0.005	0.003
firm's age						
< 1 year	0.003	0.001	0.004	0.001	0.004	0.001
1-5 years	0.001	0.001	0.001	0.001	0.001	0.001
10-20 years	0.000	0.000	0.000	0.001	0.001	0.001
> 20 years	-0.001	0.001	-0.001	0.001	-0.001	0.001
trend	-0.002	0.000	-0.003	0.000	-0.004	0.000
constant	0.107	0.014	0.148	0.020	0.141	0.025
rho	0.206	0.014	0.217	0.019	0.204	0.021
sigma	0.078	0.000	0.072	0.000	0.073	0.000
Number of observations	461247		253882		186615	
Number of clusters	64327		41663		29455	
Log-Likelihood	548021		277201		199268	

Notes: Asymptotic standard erorrs are robust to the opresence of repeated observations within firms and on the same employee.

Table 5. ML estimates of wage changes and downward wage rigidity.

Rigidity type	proportional		proportional		proportional	
Measurement error	normal		normal		normal	
Sample	all workers		stable workers		stable men	
dip. Var. : ln(wage change)	Coeff.	robust s.e.	Coeff.	robust s.e.	Coeff.	robust s.e.
ln(firm size)	0.002	0.000	0.001	0.000	0.001	0.000
ln(firm size) squared	0.000	0.000	0.000	0.000	0.000	0.000
worker's age	-1.514	0.098	-1.631	0.128	-1.660	0.147
age squared	0.014	0.001	0.014	0.002	0.015	0.002
growing firm	0.006	0.001	0.007	0.001	0.007	0.001
shrinking firm	-0.004	0.001	-0.003	0.001	-0.003	0.001
sector1	0.017	0.002	0.013	0.002	0.013	0.002
sector2	0.006	0.001	0.006	0.001	0.005	0.001
sector3	0.004	0.001	0.004	0.001	0.004	0.001
sector5	0.004	0.000	0.000	0.001	-0.001	0.001
sector6	0.008	0.001	0.002	0.002	0.001	0.002
sector7	0.009	0.001	0.005	0.001	0.004	0.001
north East	0.002	0.000	0.000	0.000	0.000	0.000
centrer	-0.002	0.000	-0.001	0.001	-0.002	0.001
south	-0.004	0.001	-0.005	0.001	-0.006	0.001
female	-0.017	0.001	-0.020	0.001		
female*age	0.281	0.028	0.356	0.036		
manager	0.027	0.001	0.026	0.001	0.026	0.001
apprentice	0.080	0.001	0.082	0.003	0.073	0.003
white-collar	0.033	0.001	0.029	0.001	0.029	0.002
white collar * age	-0.513	0.031	-0.437	0.032	-0.441	0.039
inflation	0.792	0.087	0.577	0.095	0.625	0.112
lagged inflation	-0.106	0.070	-0.199	0.049	-0.184	0.057
unemployment rate						
level	-0.039	0.012	-0.045	0.018	-0.039	0.022
first difference	-0.012	0.001	-0.014	0.001	-0.015	0.001
lagged first differ.	0.003	0.002	0.004	0.002	0.005	0.003
firm's age						
< 1 year	0.003	0.001	0.005	0.001	0.004	0.001
1-5 years	0.001	0.001	0.001	0.001	0.001	0.001
10-20 years	-0.001	0.001	0.000	0.001	0.000	0.001
> 20 years	-0.001	0.001	-0.002	0.001	-0.001	0.001
trend	-0.003	0.000	-0.004	0.000	-0.004	0.000
constant	0.109	0.015	0.151	0.022	0.143	0.027
rho	0.611	0.012	0.633	0.015	0.632	0.018
sigma(m)	0.031	0.000	0.029	0.000	0.030	0.000
sigma(e)	0.0741	0.0003	0.0685	0.0005	0.0704	0.0006
Number of observations	461247		253882		186615	
Number of clusters	64327		41663		29455	
Log-Likelihood	537709		316235		227865	

Table 6. ML estimates of wage changes and downward wage rigidity.

Rigidity type	proportional		proportional		proportional	
Measurement error	mixed		mixed		Mixed	
Sample	all workers		stable workers		stable men	
dip. Var. : ln(wage change)	Coeff.	robust s.e.	Coeff.	robust s.e.	Coeff.	robust s.e.
ln(firm size)	0.0021	0.0001	0.0013	0.0002	0.0016	0.0002
ln(firm size) squared	-0.0001	0.0000	-0.0001	0.0000	-0.0001	0.0000
worker's age	-1.1000	0.0754	-1.2785	0.1004	-1.3423	0.1202
age squared	0.0103	0.0009	0.0114	0.0012	0.0120	0.0015
growing firm	0.0036	0.0003	0.0048	0.0004	0.0052	0.0005
shrinking firm	-0.0029	0.0003	-0.0023	0.0004	-0.0024	0.0005
sector1	0.0154	0.0006	0.0137	0.0007	0.0129	0.0008
sector2	0.0049	0.0003	0.0055	0.0005	0.0046	0.0006
sector3	0.0036	0.0003	0.0048	0.0004	0.0046	0.0004
sector5	0.0029	0.0003	0.0010	0.0004	0.0005	0.0005
sector6	0.0031	0.0004	-0.0004	0.0005	-0.0011	0.0006
sector7	0.0049	0.0004	0.0034	0.0004	0.0023	0.0006
north East	0.0006	0.0002	-0.0001	0.0003	-0.0001	0.0004
centrer	-0.0018	0.0003	-0.0014	0.0003	-0.0018	0.0004
south	-0.0038	0.0003	-0.0042	0.0004	-0.0049	0.0004
female	-0.0128	0.0008	-0.0146	0.0011		
female*age	0.2069	0.0216	0.2578	0.0294		
manager	0.0196	0.0010	0.0207	0.0010	0.0208	0.0011
apprentice	0.0711	0.0011	0.0761	0.0022	0.0684	0.0027
white-collar	0.0244	0.0008	0.0222	0.0010	0.0224	0.0012
white collar * age	-0.3919	0.0198	-0.3569	0.0251	-0.3627	0.0306
inflation	0.4775	0.0195	0.4258	0.0248	0.4847	0.0305
lagged inflation	-0.1121	0.0111	-0.1791	0.0145	-0.1606	0.0178
unemployment rate						
level	-0.0323	0.0020	-0.0443	0.0025	-0.0412	0.0031
first difference	-0.0106	0.0002	-0.0118	0.0002	-0.0127	0.0003
lagged first differ.	0.0009	0.0003	0.0025	0.0004	0.0035	0.0005
firm's age					0.0022	0.0010
< 1 year	0.0023	0.0006	0.0027	0.0008		
1-5 years	0.0007	0.0004	0.0005	0.0005	0.0007	0.0006
10-20 years	-0.0002	0.0003	-0.0003	0.0004	0.0000	0.0005
> 20 years	-0.0006	0.0003	-0.0013	0.0004	-0.0010	0.0005
trend	-0.0025	0.0000	-0.0031	0.0001	-0.0032	0.0001
constant	0.1122	0.0027	0.1436	0.0036	0.1380	0.0044
rho	0.6798	0.0344	0.5680	0.0351	0.5146	0.0376
sigma(m)	0.1007	0.0003	0.0968	0.0004	0.0969	0.0005
sigma(e)	0.0426	0.0001	0.0435	0.0002	0.0456	0.0002
p	0.5948	0.0100	0.6611	0.0158	0.6546	0.0196
Number of observations	461247		253882		186615	
Number of clusters	64327		41663		29455	
Log-Likelihood	484456		290798		207893	

Table 7. ML estimates of wage changes and downward wage rigidity.

Rigidity type	threshold		threshold		threshold	
Measurement error	normal		normal		normal	
Sample	all workers		stable workers		stable men	
dip. Var. : ln(wage change)	Coeff.	robust s.e.	Coeff.	robust s.e.	Coeff.	robust s.e.
ln(firm size)	0.0019	0.0002	0.0006	0.0002	0.0010	0.0003
ln(firm size) squared	-0.0001	0.0000	0.0000	0.0000	-0.0001	0.0000
worker's age	-1.5105	0.1013	-1.6382	0.1310	-1.6544	0.1538
age squared	0.0145	0.0013	0.0144	0.0016	0.0145	0.0019
growing firm	0.0055	0.0004	0.0071	0.0005	0.0072	0.0006
shrinking firm	-0.0038	0.0004	-0.0027	0.0005	-0.0029	0.0007
sector1	0.0174	0.0008	0.0134	0.0009	0.0128	0.0010
sector2	0.0058	0.0005	0.0056	0.0006	0.0047	0.0007
sector3	0.0039	0.0004	0.0040	0.0005	0.0039	0.0006
sector5	0.0043	0.0004	0.0001	0.0005	-0.0007	0.0006
sector6	0.0078	0.0005	0.0020	0.0007	0.0012	0.0008
sector7	0.0090	0.0005	0.0054	0.0006	0.0040	0.0007
north East	0.0015	0.0003	0.0001	0.0004	0.0000	0.0005
centrer	-0.0022	0.0003	-0.0012	0.0004	-0.0018	0.0005
south	-0.0040	0.0004	-0.0048	0.0005	-0.0058	0.0006
female	-0.0173	0.0011	-0.0205	0.0015		
female*age	0.2837	0.0292	0.3596	0.0388		
manager	0.0267	0.0013	0.0264	0.0013	0.0258	0.0014
apprentice	0.0806	0.0012	0.0827	0.0022	0.0741	0.0029
white-collar	0.0334	0.0011	0.0289	0.0013	0.0287	0.0016
white collar * age	-0.5112	0.0266	-0.4358	0.0327	-0.4393	0.0388
inflation	0.7920	0.0262	0.5718	0.0324	0.6191	0.0388
lagged inflation	-0.1020	0.0151	-0.1964	0.0192	-0.1817	0.0230
unemployment rate						
level	-0.0400	0.0026	-0.0453	0.0033	-0.0399	0.0040
first difference	-0.0119	0.0002	-0.0139	0.0003	-0.0147	0.0004
lagged first differ.	0.0029	0.0004	0.0039	0.0005	0.0047	0.0006
firm's age						
< 1 year	0.0030	0.0008	0.0050	0.0010	0.0043	0.0013
1-5 years	0.0012	0.0005	0.0013	0.0006	0.0014	0.0008
10-20 years	-0.0007	0.0004	0.0002	0.0005	0.0005	0.0006
> 20 years	-0.0010	0.0004	-0.0017	0.0005	-0.0012	0.0007
trend	-0.0025	0.0001	-0.0038	0.0001	-0.0039	0.0001
constant	0.1079	0.0037	0.1510	0.0047	0.1428	0.0056
alpha	0.0688	0.0005	0.0716	0.0007	0.0728	0.0008
lamda	0.0 .		0.0 .		0.0 .	
sigma(m)	0.0335	0.0002	0.0321	0.0002	0.0330	0.0003
sigma(e)	0.0733	0.0001	0.0675	0.0001	0.0692	0.0002
Number of observations	461247		253882		186615	
Number of clusters	64327		41663		29455	
Log-Likelihood	542974		318596		229450	

Table 8: Proportional rigidity with normal measurement error. ρ depends on X.

ln(wage change)	Coeff.	Robust s.e.
worker's age	-1.3039	0.1020
age squared	0.0121	0.0013
growing firm	0.0067	0.0004
shrinking firm	-0.0032	0.0004
sector1	0.0166	0.0007
sector2	0.0060	0.0005
sector3	0.0034	0.0004
sector5	0.0036	0.0004
sector6	0.0071	0.0005
sector7	0.0082	0.0005
north East	0.0013	0.0003
centrer	-0.0021	0.0004
south	-0.0042	0.0004
female	-0.0161	0.0011
female*age	0.2533	0.0296
manager	0.0255	0.0013
apprentice	0.0780	0.0012
white-collar	0.0323	0.0011
white collar * age	-0.4896	0.0269
inflation	1.2192	0.0253
lagged inflation	0.1531	0.0145
unemployment rate		
level	-0.0835	0.0026
first difference	-0.0115	0.0003
lagged first differ.	0.0116	0.0003
firm's age		
< 1 year	0.0026	0.0008
1-5 years	0.0008	0.0005
10-20 years	0.0004	0.0004
> 20 years	-0.0020	0.0004
constant	0.1003	0.0039
sigma(m)	0.0277	0.0003
sigma(e)	0.0767	0.0001
Rho equation:		
ln_dipmed	-0.6657	0.0221
ln_dipmed2	0.0796	0.0018
1986	-1.5903	0.0709
1987	-0.7859	0.0696
1988	-0.4966	0.0806
1989	-0.9543	0.0811
1990	-0.8123	0.0924
1991	0.9879	0.0792
1992	1.4614	0.0751
1993	3.1183	0.0778
1994	1.9812	0.0810
1995	1.6698	0.0788
constant	0.9534	0.0778
Number of observations	461247	
Number of clusters	64327	
Log-Likelihood	536574	

Figure 2: Proportional Rigidity Model

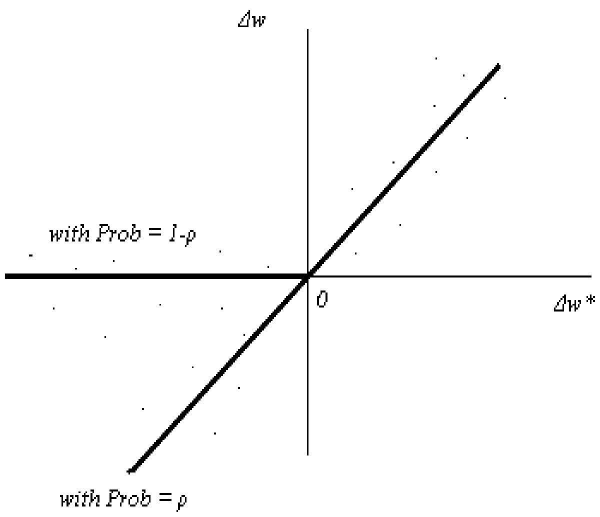


Figure 3: Threshold rigidity model

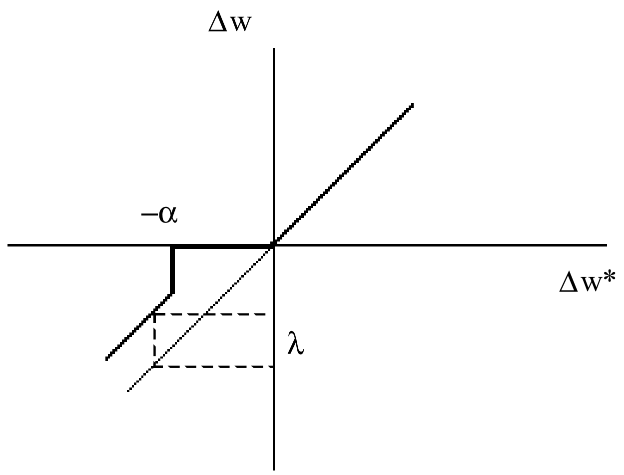


Figure 4A: Proportion of Wage Cuts, Freezes and Inflation

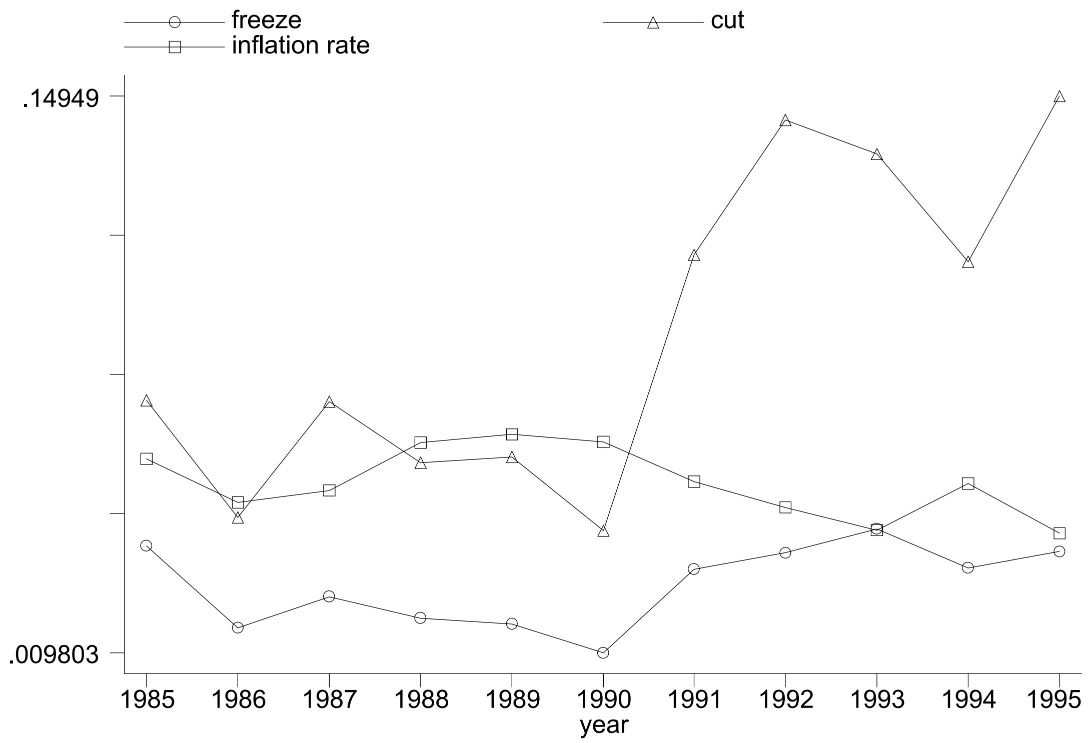


Figure 4B: Nominal Wage Changes and Inflation

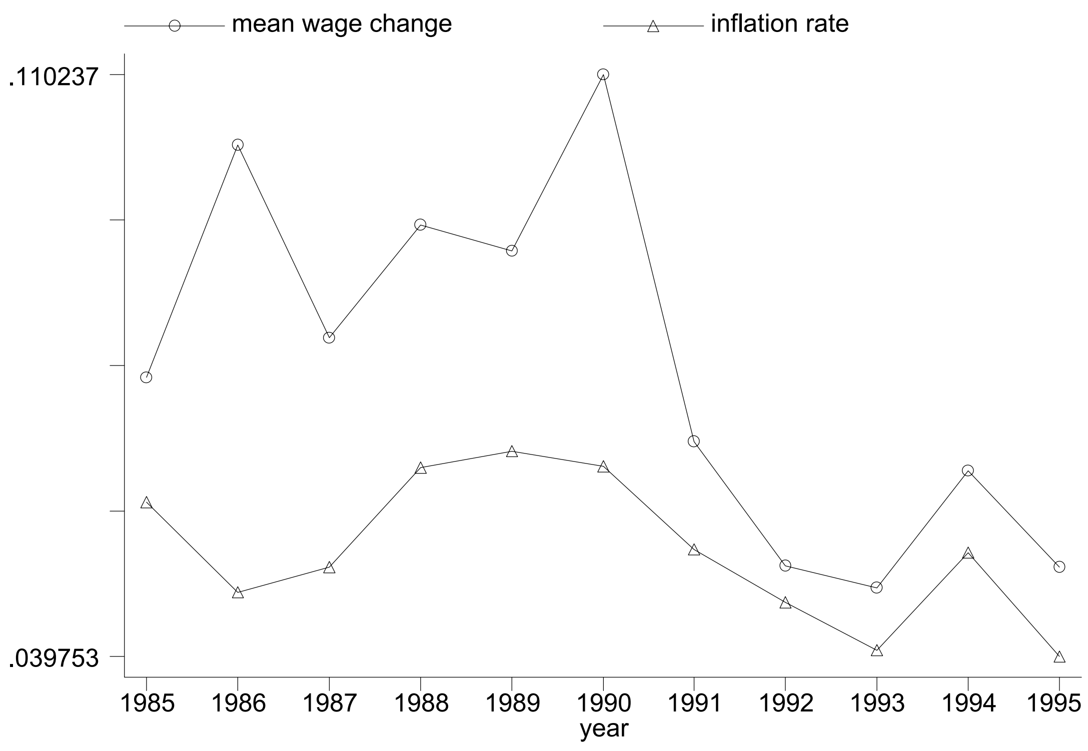


Figure 5: Effects of downward nominal wage rigidity at different rates of inflation

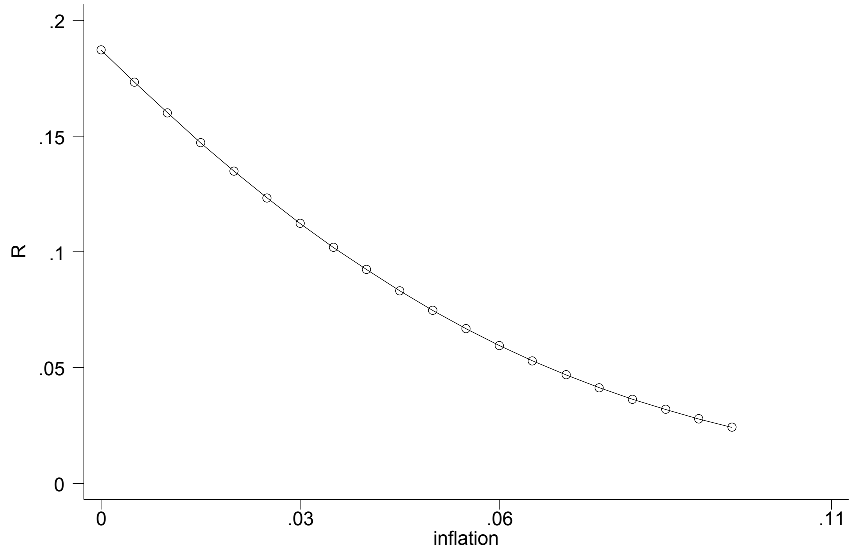


Table 9. Long-run unemployment consequences of DNWR

inflation Rate	SU	Excess long-run unemployment rate a=0.1	a=0.2	a=0.3	a=0.4	a=0.5
0.000	0.009	9.089	4.544	3.030	2.272	1.818
0.005	0.008	8.188	4.094	2.729	2.047	1.638
0.010	0.007	7.355	3.678	2.452	1.839	1.471
0.015	0.007	6.588	3.294	2.196	1.647	1.318
0.020	0.006	5.883	2.941	1.961	1.471	1.177
0.025	0.005	5.237	2.619	1.746	1.309	1.047
0.030	0.005	4.649	2.324	1.550	1.162	0.930
0.035	0.004	4.113	2.057	1.371	1.028	0.823
0.040	0.004	3.628	1.814	1.209	0.907	0.726
0.045	0.003	3.190	1.595	1.063	0.798	0.638
0.050	0.003	2.796	1.398	0.932	0.699	0.559
0.055	0.002	2.443	1.222	0.814	0.611	0.489
0.060	0.002	2.127	1.064	0.709	0.532	0.425
0.065	0.002	1.846	0.923	0.615	0.462	0.369
0.070	0.002	1.597	0.799	0.532	0.399	0.319
0.075	0.001	1.377	0.689	0.459	0.344	0.275
0.080	0.001	1.183	0.592	0.394	0.296	0.237
0.085	0.001	1.013	0.507	0.338	0.253	0.203
0.090	0.001	0.865	0.432	0.288	0.216	0.173
0.095	0.001	0.735	0.368	0.245	0.184	0.147

APPENDIX: DERIVATION OF THE MODELS' LIKELIHOOD FUNCTIONS

1. Proportional rigidity model

In this appendix I sketch the derivation of the likelihood functions of the various models that have been estimated. See Ferh and Goette (2000), Knopick (2001), Maddala (1983) for the details.

Consider the model:

$$(1) \qquad \Delta y_{it} = \begin{cases} X_{it}\beta + \varepsilon_{it} + \mu_{it} & \text{if } \; 0 \le X_{it}\beta + \varepsilon_{it} & \text{regime1} \\ \mu_{it} & \text{if } \; X_{it}\beta + \varepsilon_{it} < 0 \; \text{ and } \; D_{it} = 1 & \text{regime2} \\ X_{it}\beta + \varepsilon_{it} + \mu_{it} & \text{if } \; X_{it}\beta + \varepsilon_{it} < 0 \; \text{ and } \; D_{it} = 0 & \text{regime3} \end{cases}$$

For ease of notation, let us drop the *t* subscript and write simply c_i for $X_{it}\beta$ and y_i instead of Δy_i. The random term ε is assumed to be normally distributed with zero mean and variance σ_ε^2. When present, the measurement error term μ is assumed to be uncorrelated with ε.

No measurement error

In this case μ_{it}=0 and the likelihood function for model (1) is obtained as:

$$(2) \qquad L = \prod_{y_i>0} \frac{1}{\sigma_\varepsilon}\phi\left(\frac{y_i - c_i}{\sigma_\varepsilon}\right) \cdot \prod_{y_i=0} \rho \frac{1}{\sigma_\varepsilon}\Phi\left(\frac{-c_i}{\sigma_\varepsilon}\right) \cdot \prod_{y_i<0} (1-\rho)\frac{1}{\sigma_\varepsilon}\phi\left(\frac{y_i - c_i}{\sigma_\varepsilon}\right),$$

with the three pieces of the likelihood function corresponding to the three regimes in (1). In (2) ϕ denotes the density function of the normal standard distribution and Φ the corresponding distribution function. Note how the probability that the observation is affected (not affected) by DNWR enters in the definition of the second (third) term of the likelihood function.

Normal measurement error

To obtain the model's likelihood function we need to derive the density function of y_i, where $y = c + \varepsilon + \mu$. It is here assumed that μ is normally distributed with zero mean and variance σ_μ^2. The density function of y can be written as the sum of three joint densities, as follows:

(3) $$f_y(y) = f_1(y \text{ , regime 1}) + f_2(y \text{ , regime 2}) + f_1(y \text{ , regime 3}) =$$

$$= f_1(y \ , -c \le \varepsilon) + f_2(y \ , -c > \varepsilon \ , D{=}1) + f_3(y \ , -c > \varepsilon \ , D{=}0)$$

$$= f_1(y \ , -c \le \varepsilon) + f_D(D{=}1)\, f_{2'}(y \ , -c > \varepsilon) + f_D(D{=}0)\, f_{3'}(y, -c > \varepsilon)$$

$$= f_1(y \ , -c \le \varepsilon) + \rho\ f_{2'}(y \ , -c > \varepsilon) + (1\text{-}\rho)\, f_{3'}(y \ , -c > \varepsilon)$$

where f_D is the probability distribution of *D*, a random variable assumed to be independent of both *y* and ε.

To derive the joint density f_1, we first note that y depends on both ε and μ. Then, the joint occurrence of the events $(y \ , -c \le \varepsilon)$ can be expressed in terms of the joint distribution of ε and μ. As both random terms are assumed to be normally distributed (and uncorrelated), a standard result implies the following form for their joint distribution:

(4) $$f(\varepsilon \ , \mu) = \frac{1}{2\pi\sqrt{\sigma_\varepsilon^2 \sigma_\mu^2}} \exp\left\{ -\frac{1}{2}\left[\left(\frac{\varepsilon}{\sigma_\varepsilon}\right)^2 + \left(\frac{\mu}{\sigma_\mu}\right)^2 \right] \right\}.$$

Note that when $-c \le \varepsilon$, (1) implies that $\mu = y - c - \varepsilon$. Therefore, we can write:

(5) $$f_1(y, -c \le \varepsilon) = \frac{1}{2\pi\sqrt{\sigma_\varepsilon^2 \sigma_\mu^2}} \int_{-c}^{\infty} \exp\left\{ -\frac{1}{2}\left[\left(\frac{\varepsilon}{\sigma_\varepsilon}\right)^2 + \left(\frac{y - c - \varepsilon}{\sigma_\mu}\right)^2 \right] \right\} de.$$

(see Maddala, 1983, p. 284). The integral in (5) can be given a closed form solution after a few algebraic steps (see Knoppick, 2001), which provides the following final expression:

(6) $$f_1(y\,,\,-c \le \varepsilon\,) = \frac{1}{\sqrt{\sigma_\varepsilon^2 + \sigma_\mu^2}}\,\phi\!\left(\frac{y_i - c_i}{\sqrt{\sigma_\varepsilon + \sigma_\mu}}\right).\left(1 - \Phi\!\left(-sc - \frac{y - c_i}{s\sigma_\mu^2}\right)\right).$$

To derive $f_{2'}$, note that ε and μ are independent and therefore:

(7) $$f_{2'}(\,y\;\;,\,-c > \varepsilon\,) = f_\mu(\,y\,)\,f\,(-c > \varepsilon\,) = \frac{1}{\sigma_\mu}\,\phi\!\left(\frac{y_i}{\sigma_\mu}\right).\Phi\!\left(\frac{-c_i}{\sigma_\varepsilon}\right).$$

Finally, the derivation of $f_{3'}$ follows the same steps as for f_1. The difference is that now the integration in (5) runs from $-\infty$ to –*c*, instead to from –*c* to $+\infty$, giving:

(8) $$f_{3'}(\,y\;\;,\,-c > \varepsilon\,) = \frac{1}{\sqrt{\sigma_\varepsilon^2 + \sigma_\mu^2}}\,\phi\!\left(\frac{y_i - c_i}{\sqrt{\sigma_\mu^2 + \sigma_\varepsilon^2}}\right).\Phi\!\left(-sc - \frac{y - c_i}{s\sigma_\mu^2}\right).$$

The likelihood function for model (1) with normally distributed measurement error is therefore obtained by replacing (6)-(8) in (3), and re-interpreting the resulting density function of *y* given the vector of parameters as a function of the parameters for given y.

Mixed measurement error

If it is assumed that only a fraction (1-*p*) of the observations is affected by measurement error, while the remaining *p* is not, the definition of the observed dependent variable *y* is slightly more complex than in (1). Specifically, with probability (1-*p*), *y* is described by the system in (1) while, with probability *p*, *y* is instead given by:

$$(1') \qquad \Delta y_{it} = \begin{cases} X_{it}\beta + \varepsilon_{it} & \text{if } \ 0 \le X_{it}\beta + \varepsilon_{it} \\ 0 & \text{if } \ X_{it}\beta + \varepsilon_{it} < 0 \ \text{ and } \ D_{it} = 1 \\ X_{it}\beta + \varepsilon_{it} & \text{if } \ X_{it}\beta + \varepsilon_{it} < 0 \ \text{ and } \ D_{it} = 0 \end{cases} .$$

The individual's likelihood contribution is the expected value (taken over the discrete distribution of *D*) of the individual's contributions in the likelihoods we have derived earlier for the case of normal measurement error and no measurement error. The likelihood function obtains as the product of the individuals' contributions in the sample:

$$(9) \qquad L = \prod_{y_i} \ln\left((1-p)L_i^{nom} + pL_i^{m}\right)$$

where

$$(10) \quad L_i^{nom} = \left[\frac{1}{\sigma_\varepsilon}\phi\left(\frac{y_i - c_i}{\sigma_\varepsilon}\right)\right]^{1(y>0)} \left[\rho\frac{1}{\sigma_\varepsilon}\Phi\left(\frac{-c_i}{\sigma_\varepsilon}\right)\right]^{1(y=0)} \left[(1-\rho)\frac{1}{\sigma_\varepsilon}\phi\left(\frac{y_i - c_i}{\sigma_\varepsilon}\right)\right]^{1(y<0)},$$

is the individual's likelihood contribution in the case of *no measurement error*, and

$$(11) \quad L_i^m = [f_1]^{1(y>0)}\,[f_2]^{1(y=0)}\,[f_3]^{1(y<0)}$$

is the individual's likelihood contribution in the case of normal measurement error. The f functions in (11) are defined in (6)-(8), while 1(*j*) is an indicator function taking value 1 if expression j is true, and 0 otherwise.

2. Threshold rigidity model

For the threshold rigidity model, the dependent variable *y* is defined as:

$$(13) \qquad \Delta y_{it} = \begin{cases} X_{it}\beta + \varepsilon_{it} + \mu_{it} & \text{if } \ 0 \le X_{it}\beta + \varepsilon_{it} & \text{regime1} \\ \mu_{it} & \text{if } \ -\alpha \le X_{it}\beta + \varepsilon_{it} < 0 & \text{regime2} \\ X_{it}\beta + \varepsilon_{it} + \mu_{it} + \lambda & \text{if } \ X_{it}\beta + \varepsilon_{it} < -\alpha & \text{regime3} \end{cases}$$

The measurement error variable is assumed to be normally distributed. Once again the density distribution function can be written as:

(14) $f_y(y) = f_1(y\text{, regime 1}) + f_2(y\text{, regime 2}) + f_1(y\text{, regime 3}) =$

where now the relevant regimes are the ones in (13). Correspondingly, the density function is slightly different than the one for the proportional rigidity with normal measurement error (equations (3) and (6)-(8)) to reflect the presence of α and λ in model (13), but are easily derived:

$$(15)\quad f_y(y) = \frac{1}{\sqrt{\sigma_\varepsilon^2+\sigma_\mu^2}}\phi\left(\frac{y_i-c_i}{\sqrt{\sigma_\varepsilon+\sigma_\mu}}\right).\left(1-\Phi\left(-sc-\frac{y-c_i}{s\sigma_\mu^2}\right)\right)$$

$$+\frac{1}{\sqrt{\sigma_\varepsilon^2+\sigma_\mu^2}}\phi\left(\frac{y_i}{\sqrt{\sigma_\varepsilon+\sigma_\mu}}\right).\left(\Phi\left(\frac{-c_i}{\sigma\varepsilon}\right)-\Phi\left(\frac{y-\alpha}{\sigma\varepsilon}\right)\right)+$$

$$+\frac{1}{\sqrt{\sigma_\varepsilon^2+\sigma_\mu^2}}\phi\left(\frac{y_i-c_i-\lambda}{\sqrt{\sigma_\mu^2+\sigma_\varepsilon^2}}\right).\Phi\left(-s(c+\alpha)-\frac{y-c_i-\lambda}{s\sigma_\mu^2}\right).$$